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CYCLICAL VARIATIONS IN THE DURATION
OF UNEMPLOYMENT SPELLS

by

Miles Corak

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Miles Corak

Business and Labour Market Analysis
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EXECUTIVE SUMMARY

It is often accepted as a truism that the higher the economy-wide unemployment rate, the longer are the durations of individual unemployment spells. Aggregate data on the duration of interrupted unemployment spell lengths certainly display this pattern. Economic theory, however, does not offer unambiguous implications concerning this relationship, and recently its empirical validity has come into question. Darby, Haltiwanger, and Plant (1986) have argued that the economy-wide average duration of unemployment rises during recessions because the composition of the unemployed becomes more heavily weighted with individuals that "normally" experience long unemployment spells, not because individual spells are any longer than they normally would be.

The purpose of the research reported in this paper is to examine one implication of this view, namely that controlling for individual characteristics the probability of leaving unemployment is invariant over the course of the business cycle. While there are many existing Canadian studies that examine the duration of unemployment spells, few have adopted a long enough time horizon to shed any light on this issue. The data used are derived from the Annual Work Patterns Survey for the years 1978 to 1980 and 1982 to 1985. Accelerated life-time models under a variety of distributional assumptions are used to examine unemployment spells from each year of data.

Three major results are obtained. First, the hypothesis that individual exit probabilities are constant over the cycle is rejected, but the direction in which they mover is not clear-cut. There is strong evidence to suggest that the duration of individual unemployment spells increased between 1980 and 1982 as the economy moved into recession, and then decreased as recovery took hold. However, the results also reveal increases in spell durations when the unemployment rate was falling, most notably between 1984 and 1985.

Second, the pattern of change varies across individual characteristics, most notably across age. The results document an important and long-lasting deterioration in the relative position of older unemployed individuals. The average spell duration of the old not only increased during the onset of the 1981-82 recession, but it continued to increase during each year of the subsequent recovery. This contrasts sharply with the experience of individuals in the prime-aged, and younger age groups.

Finally, from a methodological stand point, it is found that the estimation results are robust across functional forms, but

not over time. Researchers employing parametric methods to
analyze the duration of unemployment spells should be aware that
the magnitude and signs of their estimated coefficients may be
very sensitive to the cyclical state of the economy at the time
their data was collected.

These results permit some inferences as to the nature of the
adjustment process in the labour sector. Adjustment to severe
negative shocks might reasonably be characterized as an
evolutionary process in which older workers are, in one way or
another, moved into semi-permanent or permanent retirement, while
younger workers are rehired at increasing rates. The results
also call for continued analysis of spell durations for more
recent data, and with more extensive co-variates.

CYCLICAL VARIATIONS IN THE DURATION OF UNEMPLOYMENT SPELLS

It is often accepted as a truism that the higher the economy-wide unemployment rate, the longer are the durations of individual unemployment spells. A cursory look at the aggregate data on the duration of interrupted unemployment spell lengths from the Labour Force Survey could certainly suggest as much. Economic theory, however, does not offer unambiguous predictions concerning this relationship, and recently its empirical validity has come into question. Darby, Haltiwanger, and Plant [(1986), p.1] argue, on the basis of an analysis of U.S. data, that the "main proximate determinant of changes in the unemployment rate is variations in the level and distribution of inflows into unemployment. Since the probability of leaving unemployment [that is, the duration of an unemployment spell] is primarily determined by the characteristics of those becoming unemployed and is little affected by the business cycle, outflows from unemployment and hence the actual changes in the unemployment rate are primarily determined by the inflows." In other words, the economy-wide average duration of unemployment rises during recessions because the composition of the unemployed becomes more heavily weighted with individuals that "normally" experience long unemployment spells, not because individual spells are any longer than they normally would be.

Their view is developed in the context of a search theoretic model of unemployment with rational expectations. As the

business cycle alters the rate of job offers and the nature of
the wage offer distribution, individuals respond by altering
their search intensity and reservation wage to the extent that
the duration of their unemployment spells is unchanged. This
leaves the movement in the unemployment rate, and any movements
in the overall average spell length, to be determined by the
characteristics of the individuals becoming unemployed. Darby
and his co-authors suggest that their results lend support to the
Lilien (1982) hypothesis that the dynamics of the unemployment
rate are determined by fluctuations in the natural rate of
unemployment resulting form increases in the extent of structural
adjustments.

The purpose of the research reported in this paper is to
examine one implication of this view, namely that the individual
probability of leaving unemployment is invariant over the course
of the business cycle. While there are many existing Canadian
studies that examine the duration of unemployment spells, few
have adopted a long enough time horizon to shed any light on this
issue. Beach and Kaliski (1987) is one exception, but their
major concern is with the distribution of the burden of time
spent unemployed, and as a result they do not explicitly model
the determinants of the duration of unemployment spells. Picot
and Wannell (1987) do touch upon the relationship between
unemployment durations and cyclical shocks, but their analysis is
focused upon the experience of individuals suffering permanent
lay-offs. Their finding that older workers experience very long

spells of joblessness after a plant closing is an important element in understanding the way the adjustment process works, but it is certainly less than a comprehensive look at the experience of the labour force.

The main contribution of the present research is the explicit modelling of unemployment spell durations over a seven-year period that includes both cyclical peaks and troughs. In methodology the approach is probably closest to the work of Dynarski and Sheffrin (1990) (1987) for the United States. The data used are derived from the Annual Work Patterns Survey (AWPS) for the years 1978 to 1980 and 1982 to 1985. The AWPS is a retrospective data set containing information on the labour force status of individual respondents for each month of a given year, a different sample of individuals being interviewed each January. Thus, the seven years of available data represent a series of one year windows through which to view the dynamics of individual labour force activity. For the purposes of this study an unemployment spell is defined to begin when a transition from employment to unemployment is made, and it is defined to end when a transition back to employment from either unemployment or not in the labour force occurs.

Section 1 of the paper describes the data in greater detail and offers a justification for this particular definition of an unemployment spell. The data are potentially subject to problems of truncation from both the left and the right. The former is

dealt with by considering only unemployment spells that have
begun during the year, while the latter is explicitly addressed
by the estimation procedure. The methodology underlying this
procedure and the results obtained are the subject of section 2.
Each year of data is analyzed separately by parametric models of
spell duration. The validity of the results are limited by the
use of explicit functional forms, but an attempt is made to
assess their robustness by using as many different distributions
as possible. In particular, accelerated life-time models under
the assumption that the errors take an exponential, Weibull, log-
logistic, and log-normal distribution are all estimated, as is
Cox's proportional hazards model which is a distribution free
method.

At least three major results are obtained. First, the
analysis tends to reject the hypothesis that individual exit
probabilities are constant over the cycle, but the direction in
which they move is not clear-cut. There is strong evidence to
suggest that the duration of individual unemployment spells
increased between 1980 and 1982 - when the economy-wide
unemployment rate increased from 7.5 per cent to 11.0 per cent -
and then decreased as the recovery took hold. However, the
results also reveal increases in spell durations when the
unemployment rate was falling, most notably between 1984 and
1985. Second, the pattern of change varies across individual
characteristics, most notably across age. The results document
an important and long-lasting deterioration in the relative

position of older unemployed individuals. The average spell duration of the old, those aged 55 to 69 years of age, not only increased during the onset of the 1981-1982 recession, but it continued to increase during each year of the subsequent recovery. This contrasts sharply with the experience of individuals in the prime-aged, and younger age groups. Finally, from a methodological stand point, it is found that the estimation results are robust across functional forms, but not over time. Researchers employing parametric methods to analyze the duration of unemployment spells should be aware that the magnitude and signs of their estimated coefficients may be very sensitive to the cyclical state of the economy at the time their data was collected. It is also found that the functional form assumed for the analysis does have an important bearing on measures of location: the estimate obtained for the average duration of unemployment for any given year depends very much upon functional form.

A final section summarizes these results, underscores some important caveats, and attempts, in a speculative manner, to draw some implications for the way in which the labour sector adjusted to the business cycle of the 1980s. In particular, it is suggested that the labour sector adjusted to the adverse conditions of the early 1980s in an evolutionary manner by encouraging the permanent or semi-permanent retirement of older workers from jobs in declining firms and sectors, while at the same time hiring younger workers at increasing rates in the new

and expanding sectors. The direct substitution of older workers from declining sectors to expanding sectors was of secondary importance.

1. DATA DESCRIPTION AND A PRELIMINARY ANALYSIS

The Annual Work Patterns Survey contains retrospective information on the timing and length of periods that an individual spends in unemployment (U), employment (E), and not in the labour force (N). The Survey is administered to 5 of the 6 rotation groups in the January Labour Force Survey and addresses their labour force experience over the past year. Data are available for 1977 to 1980 and 1982 to 1985. The 1977 data are not included in this study because of a change in the questionnaire. The Survey was not administered for 1981, and it was discontinued in 1986.

Individuals are assigned a labour force status for each month of the year, and a particular algorithm is used to arrange these individual monthly states into a sequence for the entire year. The algorithm, which is described in Statistics Canada (1982), serves to maximize the length of time that an individual spends in employment. The basic unit of measure is a so-called "part-month," which may essentially be thought of as one-half month.

The analysis is conducted with the micro data tapes provided

by Statistics Canada in which the individual records are survey respondents. A spell file was created from these tapes making the unit of observation an unemployment spell. Thus, information on whether or not an individual experienced multiple spells is not used. A spell is defined to begin with an E-U transition, and to continue until a transition is made back to employment. After the first period of unemployment any transitions between unemployment and not-in-the-labour-force are ignored so that spells may end with a U-E transition or with a N-E transition. Thus, the analysis applies strictly to those individuals leaving their jobs, for whatever reason.

The justifications for using this definition of spell duration are at least three in number. First, and most importantly, permitting an unemployment spell to end when an individual left the labour force would simply imply that the sequencing algorithm was being retraced. The resulting spell durations would be artificial constructions from this algorithm rather than reflections of the actual transitions undertaken. Second, the difference between unemployment and not-in-the-labour-force is very soft in this data set suggesting that there may not be a great behaviourial distinction between the two states, and that it is therefore appropriate to aggregate them into a single state. Job search at any _one_ point during the month is sufficient to lead to an unemployment classification for the _entire_ month. An individual cannot by the definitions used in the survey be both unemployed and out of the labour force

during the same month even though most individuals are not likely
to have searched for work continuously throughout the month.[1]
Further, the requirement in the constructed spell file that the
individual have an unemployment classification immediately after
the employment separation will, to some extent, have the effect
of removing those individuals not able or willing to work from
the analysis, while the lack of distinction between subsequent
periods of U and N will explicitly permit discouragement to be
reflected in the data.[2] Finally, the definition of unemployment
permits the analysis to be conducted as a two-state model. If an
unemployment spell is permitted to end with an exit to employment
or an exit into non-participation a competing risks model would
be necessary. To avoid this some researchers have aggregated
employment with not-in-the-labour-force: examples of such work
include Darby, Haltiwanger, and Plant (1986)(1985), and much of
the work of Hasan and deBroucker (1985)(1982). If the analysis
must be restricted to a two-state model it seems preferable that
unemployment and not-in-the-labour-force be aggregated together.
Kiefer (1985), Dynarski and Shefferin (1987) and some of the work
of Clark and Summers (1979) are examples of studies that have
used such an aggregation.

A number of limitations in the data should be noted. First,
the definition of unemployment is not equivalent to the
traditional search-definition used in the Labour Force Survey.
Any direct comparisons should be made cautiously, and the
interpretation of any results should recognize the broader nature

of the definition being used. The AWPS data may also be subject
to a recall bias. Statistics Canada (1982), and Beach and
Kaliski (1984) have examined the extent of this bias, and found,
among other things, that there is an under-reporting of
unemployment for periods early in the year. This bias may also
be reflected in the results. Second, the time horizon of the
analysis is restricted to one year. Any projections of spell
durations beyond this limit are therefore questionable. Third,
the co-variates available are limited to demographic,
educational, and regional variables. Potentially important
determinants of spell duration such as reason for job separation,
reservation wages, industry/occupation, and receipt of
unemployment insurance payments are not available.

Table 1 summarizes the number of completed and truncated
spells in each year of the available data. Roughly 60 per cent
of the spells are right truncated and there are, by construction,
no left truncated spells. The large percentage of incomplete
spells is due to the one year horizon of the survey. Most spells
beginning in the later part of the year will, unless they are
rather short, be in progress at the time of the survey in January
of the following year. Further, the one year horizon in
conjunction with the definition of how a spell may be initiated
makes it impossible to observe spells longer than 11.5 months.

INSERT TABLE 1

Table 2 presents the sample proportions for each of the available co-variates by year. The co-variates may be divided into four broad groups: demographic, educational, regional, and the date at which the spell started. The sample proportions are relatively stable throughout the period, but there are some notable exceptions. The fraction of household heads, of males, and of married individuals increases slightly in 1982 at the outset of the recession but returns rather quickly to pre-recession levels in the following years. The major change with respect to age concerns the fact that the proportion of spells incurred by prime aged individuals (those 25-44 years of age) increases over the years while that incurred by younger individuals (15 to 19 years of age) falls. These changes coincide with the 1982 recession. The proportion of spells due to individuals in the oldest age category stays relatively constant. This is also the case for the proportion of spells incurred by individuals that claimed to be students at some point during the year. There is, however, a tendency for a greater fraction of spells to be accounted for by the more educated over time. The proportion of spells due to those with none or elementary education falls from a high of 23.5 per cent in 1978 to a low of 15.1 per cent in 1985, while that due to those with some post-secondary education, post-secondary graduation, or some university rises.

INSERT TABLE 2

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2. METHODOLOGY AND ESTIMATION RESULTS

Accelerated failure time models are described by Cox and Oakes [(1984),62-70,85-87], and Kalbfleisch and Prentice [(1980),33-5,54-62], and have been employed in studies of unemployment spell durations by _inter alia_ Addison and Portugal (1987). Proportional hazards models have, at least since the work of Lancaster (1979), been the preferred method for such studies, but the two methods are in fact equivalent when the exponential or Weibull distributions are assumed, which is more often than not the case.[3] Briefly, it is assumed that spell duration is influenced multiplicatively by the explanatory variables, or equivalently that the natural logarithm of spell duration is linearly related to them. If t indexes spell duration then

$$t = \exp(X\beta)\, t_0^{\sigma}$$

$$\ln t = X\beta + \sigma\epsilon \qquad (1)$$

Where ϵ may be thought of as an error with probability density function $g(\ )$, σ is a scale parameter, $t_0 = \exp(\epsilon)$, a random draw from the baseline hazard, and X is a vector of co-variates the first element of which signifies a constant. The main concern in

the present analysis is to obtain estimates of β, which, with a change of sign, may also be thought of as parameters of the hazard function. When the data set consists of truncated spell lengths least squares estimation of (1) will be biased and inefficient and resort must be made to maximum likelihood. In such cases the contribution to the likelihood of a completed unemployment spell will be the density $g((\ln t - X\beta)/\sigma)$ while that of a truncated spell will be a probability given by the survivor function $G((\ln t - X\beta)/\sigma)$. The survivor function is defined as one minus the cumulative distribution function.

The above sources offer details on the estimation procedure. In what follows the estimation is conducted by year for the exponential, Weibull, log logistic, and log normal distributions. Characteristics of these distributions and their representation in log-linear models of spell duration are catalogued in Kalbfleisch and Prentice [(1980),21-301]. Cox's proportional hazards linear regression model is also employed [Cox (1972)]. This is a distribution free method based upon the ranking of spell durations.[4]

For the sake of brevity the results for each of the five models for 1980 are summarized in Table 3. The results for all distributions over all of the years are available from the author. The omitted categories in these models are given by the variables listed first in each of the groupings of Table 2, thus the reference category is: household head, male, married, 35-44 years

of age, not currently a student, some high school education, Ontario resident, with an unemployment spell that began in the third quarter of the year. The parameter estimates, whenever they prove to be significant, are robust over the distributions, the exceptions being the intercept term and the dummy variables associated with the quarter in which the spell started. This suggests that the models will yield different estimates of location. All of the other significant variables have the same sign and are within one standard deviation of each other in magnitude.[5] This result is in contrast with those obtained by Addison and Portugal (1987) in a study of unemployment durations from a U.S. survey of displaced workers drawn from the Current Population Survey, but in broad agreement with those of Dynarski and Sheffrin (1987).

INSERT TABLE 3

For the purposes of summarizing the results over the seven years under study attention may be focused upon a single distribution. On the basis of the results reported in a preliminary analysis of the empirical hazard rates the log logistic distribution is chosen.[6] Table 4 presents the results of the estimation of this model for each year. Many of the parameters are subject to change in their significance levels and even their sign. In some years individuals that were not household heads suffered shorter spells than household heads, sometimes longer spells, and sometimes spells that were not significantly different. In this

data set gender does not appear to play a significant role in
determining spell duration, except perhaps in 1978 when females
experienced longer spells and 1982 when they experienced shorter
spells. There are also no simple statements to be made about the
role of marital status except for the fact that during 1982 at the
height of the recession married individuals tended to have shorter
spell durations than their counterparts. Likelihood ratio tests
for parameter constancy across all of the sample years were
conducted and strongly rejected.[7] The exit probabilities cannot be
considered constant across this seven year period.

INSERT TABLE 4

One significant result presented in this table concerns the
importance of age in determining spell duration. Before and during
1982 both the young and the old, the 15 to 24 and the 55 to 69 year
old groups, had significantly longer unemployment spells than the
prime age group, the 35-44 year olds. The coefficients for 15-19
year olds and 20-24 year olds rivals and often is larger than that
of the 55-69 year old group. After 1982 this pattern changes. The
young no longer experience spells any different in length than the
prime age group, while the situation of the old becomes worse as
their coefficient values rise to reach the highest levels for all
age groups over all years. Indeed, this is the only coefficient,
with the exception of that for 25-34 year olds in 1983, that is
statistically significant during the post 1982 period.

In one sense the fact that the 55-69 year olds experience significantly longer spells throughout the period under study should not come as too great a surprise. The definition of unemployment that is being used would permit retirement to be captured as a part of the unemployment spell. However, this retirement comes only after the individual spent some time searching for another job after the initial job separation so that in one sense it might be thought of as the result of discouragement or involuntary in nature rather than of disability or voluntary. Further, this possibility applies mostly to the individuals approaching or older than 65 years of age, a group that represents a very small fraction of the data set. What is important is the change in the environment faced by this group over the course of the sample period. The probability of exit from unemployment for this group fell sharply vis-à-vis other groups after the 1982 recession. Finally, it should be stressed that this discussion refers only to the oldest of the age categories, the behaviour of those 45 to 54 years old, a group often cited as requiring targeted assistance, is never significantly different from the 35-44 year old group.[8]

The coefficients associated with the dummy variables indicating the quarter in which the unemployment spell began are as a group the most significant in magnitude, and indicate that spell length increases the later the start date. This appears to be so for all specifications over all years. Dynarski and Shefferin (1987) also report such a finding. The most likely explanation is

that these variables capture the recall bias in the sample. The earlier in the year that the spell began the shorter it is reported to be. This would be consistent with the analysis presented in Statistics Canada (1982), that, as noted, documents significant under-reporting of spells in during the early part of the year. It is also possible that these results capture the influence of uncontrolled heterogeneity possibly associated with industry effects, reason for job separation, and seasonality. Heterogeneity has been an important theme in the analysis of unemployment spell durations and if left uncontrolled has potentially serious consequences for the properties of the estimates and their standard errors, and biases the results toward a finding of negative duration dependence [Kiefer (1988),671-73, Lancaster (1985)(1979)]. As a result the estimates of the scale parameter may not be entirely dependable, but in spite of the fact that unobserved heterogeneity is not explicitly controlled for a certain confidence can be expressed in the parameter estimates themselves since they are similar across the functional forms, including Cox's non-parametric procedure.[9]

Such, however, is not the case for the estimates obtained of the conditional location of the distribution. Table 5 presents the estimates of the survivor function implied by the results of Table 4. These are calculated by first employing the method of Suits (1984) to transform the dummy variables so as to represent an average member of the sample rather than the arbitrary reference

category and then by making use of the functional form for the log
logistic survivor function.[10] For this distribution the median and
the mean are the same and are therefore both given by the 50th
percentile.   For the average member of the sample mean spell
duration lies between four and five months, which is considerably
longer than that reported elsewhere in the literature. Beach and
Kaliski [(1987),263], for example, report a mean duration of about
2.5 months for the 1978 AWPS.   The difference stems from the
alternative definition of an unemployment spell used in the present
study.   In fact, the estimates presented in Table 5 lie closer to
the average annual unemployment experience of individuals than to
previously reported average spell durations. Corak (1990b) reports
estimates that vary from 3.3 to 4.6 months for the average annual
unemployment experience of the unemployed.   The fact that the
present figures tend to be slightly above those estimates is
exactly as expected since the underlying definition of an
unemployment spell will in many cases cause instances of multiple
spells for an individual to be added together along with the
intervening time spent out of the labour force. However, similar
estimates derived from the results of the Weibull distribution are
even longer since this distribution has a much greater part of the
density in the outer tail.   For example, the estimate for the
median for 1978 is 5.12 months, almost a month longer than the
estimate of 4.2 months of the log logistic distribution.[11] Thus,
the estimates of Table 5 are not robust across functional forms.

INSERT TABLE 5

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Even so, the movements that they display over the period, which is the main concern of this study, may be valid. Also included in the table is the economy-wide unemployment rate. The estimates in Table 5 offer some evidence that, controlling for the characteristics of individuals, suggests spell durations move counter-cyclically. This is certainly the case as the economy moved into recession from 1980 to 1982, but not as clear-cut during the recovery. Mean spell duration is roughly constant during 1978 to 1980, but spells that began in 1982 are estimated to be roughly one half-month longer than those that began in 1980. This figure then falls with the course of recovery, with the significant exception of 1985 when it increases again. The results presented in Table 4 reveal that many of the co-variates lose significance for this year.

The survivor functions for 15-19, 35-44, and 55-69 year old are also presented. These results control for all other characteristics of these groups. The counter-cyclical movement in average spell durations is evident for the two younger groups, but not for the older age category. For this category average spell durations increased between 1980 and 1982, but also for every year afterward. This is an attendant consequence of the movement in the parameters reported in Table 4.

Finally, it should be noted that all of these patterns appear to hold for the entire survivor function, not just the 50th

percentile. Overall about 25 per cent of an incoming cohort of unemployed individuals can be expected to spend about one year unemployed before finding another job, while 10 per cent may be expected to stay unemployed for 28 to 36 months. This latter figure, however, is questionable. The horizon of the data is limited to one year and it is therefore not strictly valid to entertain projections outside of this sample limit. Even so the movement in the entire distribution and particularly its outer tail is, like the mean spell duration, counter-cyclical. The oldest group is also an exception to this pattern.

3. CONCLUSION

The present study summarizes the results of research on the determinants of unemployment spell durations of individuals experiencing job separations in each year between 1978 to 1980 and 1982 to 1985. Accelerated failure time models that incorporate explicit assumptions concerning the functional form of the baseline hazard are estimated for each year, and for a variety of functional forms. Cox's proportional hazards procedure which does not involve parametric assumptions is also employed. The main contribution of the paper lies in the use of seven years of data that cover a period of both positive and negative shocks. Most previous studies of spell durations are restricted to limited points on the business cycle. The results obtained are robust to the functional form assumed, but not necessarily over the year of data used. It is found that the average duration of an unemployment spell increased

significantly as the economy moved into recession during the early 1980s, and that, for the most part, it decreased during the following recovery. However, even though the aggregate unemployment rate fell between 1984 and 1985 the average duration of an unemployment spell increased. There is some tendency for spell durations to move counter-cyclically, but this should in no way to be understood as an iron-clad law. Indeed, it is also observed that for some demographic groups economic recovery does little to reduce unemployment spell durations. Most notable in this regard is the change in the relationship between age and spell duration. Before the 1982 recession both the young and the old were subject to longer unemployment spells than the prime aged groups. During the subsequent recovery age no longer played a role for the young but it became more significant for the old. The spell durations of individuals aged 55 to 69 not only increased at the onset of the recession, but also during each year of the subsequent recovery. Thus, the often stated assertion that older workers have a low incidence of unemployment, but a longer duration is too simplistic: their spells are longer and increasingly so over time.

This result permits some inferences as to the nature of the adjustment process in the labour sector. Adjustment to severe negative shocks might reasonably be characterized as an evolutionary process in which older workers have, in one way or another, been moved into semi-permanent or permanent retirement, while younger workers have been rehired at increasing rates. More

detailed enquiries are required. In particular, there is a need for an examination of the retirement decision, and an assessment of the search behaviour of the old. Murphy and Topel (1987) have examined this line of inquiry for the United States, but there have been no similar analyses of Canadian data.

The present results also call for continued analysis of spell durations for more recent data, and with more extensive co-variates. The observation that the parameter estimates are not robust over time may reflect the possibility that important co-variates are omitted from the estimation. Reason for separation, unemployment insurance eligibility, industry, and occupation all suggest themselves as important influences of spell durations. The main conclusion at this study, that unemployment spell durations are not constant over the cycle should be tempered by the possibility that individual characteristics are not completely controlled, and hence that omitted variables are the cause of the instability in parameter estimates.

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ENDNOTES

1.   The possibility that the distinction between U and N is soft
is also supported by the fact that gross flows data from the Labour
Force Survey, which uses a definition of unemployment broadly
similar to the AWPS, reveal significant flows from N to E. If one
subscribed strictly to a search theoretic interpretation of
unemployment such movements are not theoretically possible. Their
prevalence in reality suggests that there is little value in
restricting the definition of unemployment in these terms.

2.   In effect this definition of unemployment accepts the claim by
Clark and Summers (1979) that many of the transitions between U and
N are spurious and do not reflect changes in behaviour. It should
be noted, however, that Flinn and Heckman (1983) challenge this
view.

3.   Cox and Oakes [(1984),71] prove that with constant explanatory
variables accelerated failure time models and proportional hazards
models coincide under the assumption of a Weibull distribution for
the baseline hazard.

4.   The parameters were established by maximum likelihood using a
Newton-Raphson algorithm by the S.A.S. version 5 procedure LIFEREG.
The sampling weights associated with the data were not used in the
estimation.   The standard errors are given by the inverse of the
observed information matrix.   In actual fact the estimation

procedure was much broader than indicated in the text and included the estimation of fully interactive models by gender for each of the distributions. The likelihood values for these models were not substantially greater with the result that the Akaike Information Criterion suggested the choice of the models with no interactions. An attempt was also made to estimate fully interactive models by province but convergence was never attained regardless of the year or distribution used. Finally, the gamma and generalized gamma distributions were also employed but convergence was much more difficult and for the majority of years under study not attained in spite of the fact that several different starting values for the estimates were used. Lawless (1982) suggests that such difficulties are often the case because of the flatness of the likelihood function. The results for these distributions are not reported in the present paper. Cox's proportional hazards model is estimated by the S.A.S. procedure PHGLM.

5. It should be noted that the important significance test concerning the scale parameter in the Weibull and log logistic is whether it differs from one, since a Weibull model with scale equal to one reduces to an exponential and a log logistic model with a scale greater than one indicates that the hazard is non-monotonic, first rising and then falling. Finally, for the Weibull model the scale parameter, σ, is equivalent to $1/\alpha$ of earlier parlance so that the results suggest that the hazard is declining with spell duration.

6.   See Corak (1990) and the related appendix, which is available from the author.

7.   The proper way to proceed in this regard is to obtain the log likelihood value from an estimation of the model over all the available data without distinguishing the year.   This forms the restricted model while the models estimated for each year separately, which together are equivalent to a fully interactive model by year, form the unrestricted model.   Unfortunately, convergence could not be attained for a model estimated on the aggregate data for the exponential distribution, which is the simplest distribution of those entertained.   This avenue was not therefore pursued.   Each year of data was used to estimate a log logistic model in which the parameters were restricted to those of 1980 and a likelihood ratio test was conducted in which the unrestricted models are given by the results of Table 6.   The likelihood ratio test statistics for 1978 through 1985 were: 104, 116, 232, 155, 184, 151.   This statistic is distributed as χ^2 with 28 degrees of freedom.   The critical χ^2 value at 1 per cent significance is 48.3, thus the null that the parameters are the same as those for 1980 may be rejected.   A more complete analysis might consider employing the estimates of each year in turn as the basis for the null hypothesis.

8.   These results also suggest that researchers should pay
particular attention to the manner in which the age variable is
entered into their models.   If actual data on age, as opposed to
categories,   is   available   a   non-linear   function   such   as   the
quadratic or a step function is preferable.

9.   Lancaster   [(1985),   162-3]   shows   that   the   estimates   of
elasticities   of   mean   spell   duration   are   correct   even   if
heterogeneity is not controlled for.   Strictly speaking, however,
his   result   applies   to   the   Weibull   model   in   which   there   are   no
truncated data.

10.   Using   the   earlier   notation,   since $G(t) = 1/(1+\gamma t^{\alpha})$ is   the
survivor function for the log logistic distribution then t may be
found for given values of $G(t)$ by recognizing that $\gamma = \exp(-X\beta)$ and
$\alpha = 1/\sigma$ where β is given by the estimated parameters of Table 7 and
σ by the estimate for the scale parameter. The calculations from
this procedure and from the transformations of Suits (1984) are
available from the author.

11.   The Weibull estimate of the mean is even longer since the mean
and the median do not coincide for this distribution, the mean
being the 63.2 percentile.

Table 1

Annual Work Patterns Survey Spell Data: 1978-80, 1982-85

Year	Total Number of Spells	Completed Spells		Truncated Spells	
1978	9 627	3 722	(38.7)	5 905	(61.3)
1979	9 522	3 592	(37.7)	5 930	(62.3)
1980	9 877	3 843	(38.9)	6 034	(61.1)
1982	14 265	5 789	(40.6)	8 476	(59.4)
1983	12 156	4 857	(40.0)	7 299	(60.0)
1984	12 168	4 593	(37.8)	7 575	(62.2)
1985	11 371	4 483	(39.4)	6 888	(60.6)

( )   indicates row per cent.

Table 2

Sample Proportions of Co-Variates: AWPS Spell Data, 1978-80 1982-85

		1978	1979	1980	1982	1983	1984	1985
1)	Demographic							
	Household head	50.2	51.0	51.5	52.8	51.1	50.5	50.7
	Not head	49.8	49.0	48.5	47.2	48.9	49.5	49.3
	Male	63.9	63.3	65.2	66.8	64.8	63.0	61.9
	Female	36.1	36.7	34.8	33.2	35.2	37.0	38.1
	Married	58.6	56.7	55.4	59.0	56.2	56.9	56.0
	Single	36.5	38.0	39.0	35.4	37.7	36.8	37.2
	Other	4.9	5.3	5.6	5.6	6.1	6.3	6.8
	35-44 years	14.6	14.6	15.0	17.1	16.5	17.6	17.5
	15-19 years	18.6	18.3	18.4	13.7	13.9	11.3	12.6
	20-24 years	22.3	22.8	23.0	22.1	23.3	22.9	22.5
	24-34 years	27.5	27.5	27.6	29.5	29.8	32.2	31.7
	45-54 years	10.6	10.5	10.1	11.4	10.2	10.2	9.4
	55-69 years	6.4	6.3	5.9	6.2	6.3	5.8	6.3
2)	Educational							
	Not student	86.4	87.2	87.5	88.7	87.9	88.2	87.5
	Student	13.6	12.8	12.5	11.3	12.1	11.8	12.5
	Some high school	57.2	59.3	58.6	59.1	58.7	58.0	58.8
	None or elementary	23.5	21.9	21.6	18.8	17.1	15.5	15.1
	Some postsecondary	6.9	6.5	6.9	7.7	8.6	9.2	8.9
	Postsecondary grad	7.8	7.9	8.3	9.7	10.4	11.1	11.4
	University	4.6	4.4	4.6	4.7	5.2	6.2	5.8
3)	Regional							
	Ontario	17.7	18.6	19.4	19.4	17.4	16.6	15.7
	Newfoundland	8.8	7.5	7.2	6.6	7.8	7.4	8.2
	P.E.I.	3.5	3.6	3.1	3.2	3.7	3.2	3.1
	Nova Scotia	7.3	6.8	6.9	6.8	7.1	7.3	7.6
	New Brunswick	9.3	9.8	9.9	8.5	9.9	9.5	9.2
	Quebec	17.3	17.5	17.0	15.9	15.2	15.3	16.5
	Manitoba	6.6	6.1	6.4	6.7	6.5	5.4	5.5
	Saskatchewan	5.8	5.6	6.3	5.8	7.2	7.4	7.5
	Alberta	9.6	10.4	10.5	11.9	11.8	15.2	14.1
	B.C.	14.1	14.1	13.3	15.2	13.4	12.7	12.6
4)	Start Date							
	3rd quarter	30.7	28.4	30.0	32.8	32.2	31.7	32.0
	1st quarter	12.0	11.5	12.3	12.0	12.5	11.8	11.8
	2nd quarter	19.0	19.1	20.7	21.3	20.1	19.2	20.8
	4th quarter	38.3	41.0	37.0	33.9	35.2	37.3	35.4

Table 3

Maximum Likelihood Estimates: AWPS Spell Data, 1980

			Functional Form		
Model	Exponential	Weibull	Log-Logistic	Log Normal	Proportional Hazards
Intercept	1.933* (0.06209)	2.030* (0.07243)	1.499* (0.07269)	1.556* (0.06957)	
Not head	0.1296* (0.04174)	0.1453* (0.04832)	0.1898* (0.04989)	0.1652* (0.04681)	-0.1203* (0.04191)
Female	0.04472 (0.03999)	0.04834 (0.04625)	0.05021 (0.04741)	0.04267 (0.04438)	-0.03416 (0.04012)
Single	-0.1188* (0.04497)	-0.1319* (0.05200)	-0.1137* (0.05371)	-0.09098* (0.05025)	0.08933* (0.04505)
Other marital	-0.005861 (0.07135)	-0.005751 (0.08254)	0.02791 (0.08578)	0.02287 (0.08118)	-0.0005781 (0.07162)
15-19 years	0.2368* (0.07475)	0.2676* (0.08653)	0.2895* (0.08892)	0.2413* (0.08352)	-0.2053* (0.07501)
20-24 years	0.2247* (0.05988)	0.2483* (0.06927)	0.2258* (0.07155)	0.1787* (0.06680)	-0.1704* (0.05996)
25-34 years	0.04584 (0.05114)	0.05359 (0.05914)	0.07983 (0.06161)	0.05980 (0.05815)	-0.04801 (0.05122)
45-54 years	-0.03448 (0.06479)	-0.03471 (0.07488)	0.01219 (0.07722)	0.004673 (0.07326)	0.01165 (0.06474)
55-69 years	0.1587* (0.08074)	0.1742* (0.09337)	0.1762* (0.09598)	0.1513* (0.08954)	-0.1300* (0.08082)
Student	-0.1701* (0.05893)	-0.1879* (0.06824)	-0.2263* (0.06993)	-0.1974* (0.06595)	0.1314* (0.05924)
Elementary	-0.03841 (0.04550)	-0.04163 (0.05263)	-0.05819 (0.05393)	-0.05563 (0.05062)	0.02592 (0.04558)
Postsecondary	0.004935 (0.06524)	0.006982 (0.07543)	-0.02360 (0.07812)	-0.02117 (0.07296)	-0.002043 (0.06528)
Postsec grad	-0.1547* (0.05824)	-0.1700* (0.06735)	-0.1459* (0.06982)	-0.1283* (0.06658)	0.1133* (0.05826)
University	0.003110 (0.07542)	0.0007567 (0.08720)	-0.05768 (0.09169)	-0.04973 (0.08579)	0.01211 (0.07545)

Table 3   (concluded)

Maximum Likelihood Estimates: AWPS Spell Data, 1980

Model	Exponential	Weibull	Log-Logistic	Log Normal	Proportional Hazards
Newfoundland	0.1945* (0.07104)	0.2089* (0.08213)	0.1877* (0.08436)	-0.1624* (0.07847)	-0.1460 (0.07103)
P.E.I.	0.1438 (0.1020)	0.1527 (0.1180)	0.1292 (0.1200)	0.1180 (0.1117)	-0.1138 (0.1021)
Nova Scotia	0.2703* (0.07471)	0.2970* (0.08641)	0.3062* (0.08748)	0.2612* (0.08117)	-0.2151* (0.07473)
New Brunswick	0.03548 (0.06190)	0.03522 (0.07155)	0.04092 (0.07351)	0.03404 (0.06946)	-0.03182 (0.06186)
Quebec	0.4137* (0.05634)	0.4598* (0.06528)	0.4954* (0.06572)	0.4329* (0.06135)	-0.3464* (0.05634)
Manitoba	0.01701 (0.07207)	0.01381 (0.08331)	-0.001669 (0.08601)	-0.0006485 (0.08103)	0.0005766 (0.07205)
Saskatchewan	0.02256 (0.07540)	0.02380 (0.08717)	0.03412 (0.08830)	0.03811 (0.08344)	-0.02231 (0.07543)
Alberta	0.01390 (0.05314)	0.01239 (0.06973)	0.003927 (0.07198)	-0.001020 (0.06791)	0.005222 (0.06034)
B.C.	-0.04343 (0.05459)	-0.05284 (0.06311)	-0.07245 (0.06569)	-0.06799 (0.06223)	0.04784 (0.05458)
1st quarter	-0.2849* (0.04563)	-0.4096* (0.05421)	-0.6558* (0.05720)	-0.5720* (0.05448)	0.5764* (0.04630)
2nd quarter	-0.2394* (0.04074)	-0.3310* (0.04799)	-0.4640* (0.04924)	-0.4333* (0.04684)	0.4448* (0.04112)
4th quarter	0.2021* (0.05389)	0.3721* (0.06476)	0.3063* (0.05622)	0.2961* (0.05120)	-0.5691* (0.05468)
Scale parameter	1.0	1.156 (0.01614)	0.8859 (0.01194)	1.457 (0.01827)	
In likelihood	-10 364.8	-10 307.5	-9 966.73	-9 752.94	

(.)   indicates standard error.
*   indicates that parameter differs from zero with at least 10 per cent significance.

Table 4

Maximum Likelihood Estimates: Log Logistic Model, 1978-80 1982-85

Model	1978	1979	1980	1982	1983	1984	1985
Intercept	1.607*	1.664*	1.499*	1.524*	1.591*	1.844*	1.747*
	(0.07471)	(0.07783)	(0.07269)	(0.05853)	(0.06656)	(0.06877)	(0.07230)
Not head	-0.1173*	-0.03724	0.1898*	0.1886*	0.1751*	0.01406	0.03627
	(0.04989)	(0.05111)	(0.04989)	(0.04194)	(0.04453)	(0.04363)	(0.04503)
Female	0.1475*	0.02967	0.05021	-0.07841*	-0.03649	0.04091	0.04648
	(0.04715)	(0.04825)	(0.04741)	(0.04057)	(0.04205)	(0.04196)	(0.04300)
Single	-0.01811	0.03956	-0.1137*	0.1487*	0.01440	0.04284	0.04784
	(0.05427)	(0.05531)	(0.05371)	(0.04557)	(0.04855)	(0.04823)	(0.05005)
Other marital	-0.02837	0.1579*	0.02791	0.1245*	0.1758*	-0.1358*	0.08942
	(0.09031)	(0.09156)	(0.08578)	(0.07095)	(0.07599)	(0.07365)	(0.07466)
15-19 years	0.2524*	0.05216	0.2895*	0.1349*	0.03012	0.03680	-0.08601
	(0.08810)	(0.09141)	(0.08892)	(0.07593)	(0.08440)	(0.08685)	(0.08825)
20-24 years	0.2246*	0.1376*	0.2258*	0.07352	0.07084	0.07224	0.007955
	(0.07096)	(0.07329)	(0.07155)	(0.0709)	(0.06341)	(0.06336)	(0.06484)
25-34 years	0.04289	0.06628	0.07983	-0.01093	0.1317*	0.03328	0.05798
	(0.06165)	(0.06455)	(0.06161)	(0.04787)	(0.05281)	(0.05234)	(0.05395)
45-54 years	-0.02613	0.002554	0.01219	-0.05294	0.05730	-0.1083	0.001885
	(0.07673)	(0.07763)	(0.07722)	(0.05924)	(0.06763)	(0.06814)	(0.07291)
55-69 years	0.1762*	0.1823*	0.1762*	0.1406*	0.2911*	0.3143*	0.2376*
	(0.09278)	(0.09589)	(0.09598)	(0.07475)	(0.08283)	(0.08875)	(0.08532)
Student	-0.1935*	-0.1369*	-0.2263*	-0.2755*	-0.09540	-0.1984*	0.04347
	(0.06764)	(0.06994)	(0.06993)	(0.05970)	(0.06401)	(0.06417)	(0.06852)
None or elem	0.06431	-0.03977	-0.05819	-0.1004*	-0.05882	0.03773	0.06243
	(0.05205)	(0.05469)	(0.05393)	(0.04540)	(0.05148)	(0.05483)	(0.05707)
Some postsec	0.01356	-0.04100	-0.02360	0.1461*	-0.1483*	-0.03597	-0.04780
	(0.07515)	(0.07932)	(0.07812)	(0.06176)	(0.06173)	(0.06181)	(0.06494)
Postsec grad	-0.003029	0.08298	-0.1459*	-0.08538	-0.1661*	-0.1060*	-0.07444
	(0.07204)	(0.07602)	(0.06982)	(0.05412)	(0.05684)	(0.05670)	(0.05822)
University	-0.08036	0.08504	-0.05768	0.1744*	-0.16997*	-0.06818	-0.05280
	(0.09048)	(0.09557)	(0.09169)	(0.07698)	(0.07683)	(0.07309)	(0.07689)

Table 4 (concluded)

Maximum Likelihood Estimates: Log Logistic Model, 1978-80 1982-85

Model	1978	1979	1980	1982	1983	1984	1985
Newfoundland	0.3673*	0.3282*	0.1877*	0.3235*	0.06881	0.2798*	-0.04625
	(0.08199)	(0.08901)	(0.08436)	(0.07527)	(0.07569)	(0.08178)	(0.07916)
P.E.I.	-0.09498	0.06242	0.1292	-0.006061	-0.03612	0.02187	0.07766
	(0.1073)	(0.1121)	(0.1200)	(0.09519)	(0.1004)	(0.1065)	(0.1144)
Nova Scotia	0.1597*	0.08813	0.3062*	0.1220*	-0.05432	-0.04829	-0.05082
	(0.08356)	(0.08825)	(0.08748)	(0.07083)	(0.7720)	(0.07814)	(0.07885)
New Brunswick	0.1441	0.06534	0.04092	-0.07926	-0.2526*	-0.1794	-0.04815
	(0.07836)	(0.07624)	(0.07351)	(0.06359)	(0.06682)	(0.06964)	(0.07423)
Quebec	0.2874*	0.3253*	0.4954*	0.1485*	0.1964*	0.1720*	0.1554*
	(0.06481)	(0.06650)	(0.06573)	(0.05368)	(0.06159)	(0.06373)	(0.06398)
Manitoba	-0.1264	-0.06674	-0.001669	0.05103	-0.05797	0.08482	-0.02158
	(0.08456)	(0.08815)	(0.08601)	(0.07074)	(0.07872)	(0.08794)	(0.08814)
Saskatchewan	-0.02554	-0.1904	0.03412	0.1330*	-0.07496	-0.08029	0.05643
	(0.09125)	(0.09187)	(0.08830)	(0.7666)	(0.07572)	(0.07710)	(0.08012)
Alberta	-0.07286	-0.07142	0.003927	0.04433	0.1079*	-0.03890	-0.007412
	(0.07438)	(0.07378)	(0.07198)	(0.05806)	(0.06511)	(0.06174)	(0.06522)
B. C.	-0.1783*	-0.05206	-0.07245	-0.1195*	-0.1558*	-0.2491*	-0.08841
	(0.06430)	(0.06721)	(0.06569)	(0.05219)	(0.06092)	(0.06289)	(0.06663)
1st quarter	-0.7297*	-0.8521*	-0.6558*	-0.4391*	-0.5078*	-0.6058*	-0.7645*
	(0.05698)	(0.05959)	(0.05720)	(0.04798)	(0.05131)	(0.05316)	(0.05427)
2nd quarter	-0.3695*	-0.5466*	-0.4640*	-0.2405*	-0.2561*	-0.4311*	-0.5358*
	(0.04966)	(0.05140)	(0.04924)	(0.04006)	(0.04444)	(0.04560)	(0.04601)
4th quarter	0.1904*	0.1586*	0.3063*	0.2621*	0.3256*	0.08533*	0.2165*
	(0.05308)	(0.05423)	(0.05622)	(0.04437)	(0.04886)	(0.04820)	(0.05113)
Scale	0.8648	0.8735	0.8858	0.9005	0.8994	0.8950	0.9014
	(0.01187)	(0.01217)	(0.01194)	(0.009967)	(0.01086)	(0.01109)	(0.01132)
In likelihood	-9 628.45	-9 352.76	-9 966.73	-15 251	-12 781.3	-12 234.2	-11 811.5

(.) indicates standard error.

* indicates that parameter differs from zero with at least 10 per cent significance.

Table 5

Estimated Log Logistic Survivor functions: 1978-80 1982-85,
Sample Average and by Selected Age Groups

Survival Quantile	1978	1979	1980	1982	1983	1984	1985
1) Unemployment Rate							
	8.3	7.4	7.5	11.0	11.9	11.2	10.5
2) Sample Average							
0.90	0.63	0.65	0.62	0.66	0.64	0.61	0.68
0.75	1.63	1.70	1.63	1.78	1.72	1.63	1.84
0.50	4.22	4.43	4.31	4.78	4.61	4.37	4.96
0.25	10.90	11.60	11.40	12.90	12.40	11.70	13.40
0.10	28.20	30.20	30.20	34.60	33.30	31.20	36.00
3) 15-19 Years							
0.90	0.73	0.64	0.72	0.72	0.60	0.60	0.61
0.75	1.88	1.66	1.91	1.94	1.61	1.60	1.63
0.50	4.85	4.33	5.05	5.22	4.32	4.28	4.39
0.25	12.60	11.30	13.40	14.00	11.60	11.40	11.80
0.10	32.50	29.50	35.30	37.70	31.10	30.60	31.80
4) 35-44 Years							
0.90	0.56	0.60	0.54	0.63	0.58	0.58	0.66
0.75	1.46	1.58	1.43	1.70	1.56	1.54	1.78
0.50	3.77	4.11	3.78	4.56	4.19	4.12	4.78
0.25	9.75	10.70	10.00	12.30	11.30	11.00	12.90
0.10	25.20	28.00	26.50	33.00	30.20	29.50	34.70
5) 55-69 Years							
0.90	0.67	0.72	0.64	0.73	0.78	0.79	0.84
0.75	1.74	1.89	1.70	1.95	2.09	2.11	2.25
0.50	4.50	4.94	4.51	5.25	5.60	5.64	6.07
0.25	11.60	12.90	11.90	14.10	15.10	15.10	16.30
0.10	30.10	33.70	31.60	38.00	40.40	40.30	44.00

Table entries for panels 2) - 5) are months of unemployment.

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